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Tse-min Lin

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Ever since the publication of *Voting* (Berelson et al. 1954), individual-level electoral studies have underscored the predictive utility of attitudes on voting behavior. More recently, developments in social psychology have led political scientists to examine closely the process by which attitudes guide behavior (Fazio 1986; Aldrich et al. 1989). Findings in this research generally suggest that attitudes must be available and accessible if they are to exert any influence on perceptions, judgments, or behavior.¹

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These two conditions for the predictive utility of attitudes raise questions about extending contemporary survey research findings to historical analyses based on aggregate data. Aggregate-level analyses usually assume an individual voter in possession of certain attitudes or “perspectives.” The relationship of these attitudes to behavior is then shown or supposed to maintain at the aggregate level. While the availability and accessibility of these attitudes may have been established by contemporary survey research, the question of whether these conditions obtain in historical contexts is seldom addressed.

Indeed, as early as 1965, the same year Samuel Hays (1965) called for a social analysis of American political history, Walter Dean Burnham (1965) cautioned against examining historical aggregate data in the light of relationships established between findings of survey research and contemporary aggregate data. In the same spirit, Burnham (1974: 1019) demonstrated that the historical level of political cognition in mass electorates was not more or less constant and low, as has been established in contemporary survey research, but was subject to great fluctuations that were “heavily dependent upon historical and sociological contexts.” Echoing Hays’s call, Burnham advocated a “comparative sociology” of past voting behavior, which, he argued, should be more conducive to the development of theories of democratic linkage and of American politics.

This article builds on theories of American political development to examine the historical significance of economic voting. Aggregate-level studies of economic voting typically assume a self-interested voter who is informed about the economic performance of the incumbent party and makes his or her vote choice on the basis of such information.² The assumption has been tested, with some success, in cross-sectional analyses based on contemporary survey research.³ Many aggregate-level studies, however, estimate regression models with constant parameters over a long historical period, implying an assumption of temporal homogeneity on individual voting behavior. Following the critique of Larry W. Isaac and Larry J. Griffin (1989) and Griffin and Isaac (1992), I argue that such an implicit assumption is “ahistorical.” By testing the stability of regression coefficients and estimating models with time-varying parameters, I demonstrate that the statistical significance of economic voting has varied with the changing shape of the American political universe.

The Changing Political Universe

In "The Changing Shape of the American Political Universe," Burnham (1965) pointed out the fundamental differences between the political "universe" of voting in the late nineteenth century and the contemporary one taken for granted by many researchers. Among the markedly different characteristics are a solid "core" electorate, the extremely low salience of class antagonism in voting, the influence of rurality on the intensity and uniformity of voting participation, and relatively stable attachment to the political parties. The "system of 1896" led to the eventual collapse of this essentially preindustrial democratic system. After 1900, the old political order eroded as turnout declined, the core electorate shrank, split-ticket voting increased, and the amplitude of partisan swing became substantially larger.

This shift in the shape of the American political universe is illuminated by the "community-society" dimension proposed by Hays (1975) as a conceptual framework for understanding American political history. The pre-urban-industrial human relationship was characterized by "personal, community, face-to-face contacts." At this end of the continuum, family, school, and church were the social institutions of primary importance, giving rise to political perspectives focusing on ethnocultural and ethnoreligious matters. People perceived their political world in terms of community-oriented scope and variety, and the preferred mechanism of decision making was situated at the local level. The force of modernization in the late nineteenth century, however, brought about a shift toward "impersonal, mass relationships in the wider society." Functional and economic institutions gained eminence in the social structure as political perspectives became more cosmopolitan, and preferences on the level of decision making moved upward.

Hays's work pioneered a school of ethnocultural and ethnoreligious political analysts who contended that, until the turn of the century, the most important determinants of party loyalties and voting behavior were ethnic and religious concerns rather than economic and class issues. According to Richard McCormick's (1974) summary of the theories of ethnocultural voting, political affiliation in the American past is explained by (1) negative reference group feelings, (2) differences in religious beliefs and worldviews, and (3) efforts to extend or protect cultural practices. These tendencies or (in the words of Hays) "impulses" are more symbolic than rational or goal-oriented.

Although McCormick notes the reluctance of some theorists to go beyond the Progressive view of socioeconomic voting, he correctly points out that the “ethnocultural approach to political analysis embodies a denial that voters respond wholly rationally to different party policies on economic questions” (1974: 374). The political significance of an ethnocultural rather than a socioeconomic base of partisan identifications has many systemic implications for the third-party system, the most conspicuous of which is the high participation and long-term stability that Burnham so lucidly demonstrates (Kleppner 1981).

It is noteworthy that class-based socioeconomic voting is not exactly the same as economic voting. Socioeconomic voting implies voting for or against *a particular party* based on the party’s class-based economic policies, whereas economic voting implies voting for or against *the incumbent party* based on the conditions of the economy. When the economy is bad, however, widespread socioeconomic voting by financially disadvantaged social groups would, in aggregate, look just like economic voting. In any case, it is clear that ethnocultural voting is as antithetical to economic voting as it is to socioeconomic voting. A voter situated in Hays’s “community” did not enjoy as much psychological freedom of making an economically based voting decision as the cosmopolitan voter living in an impersonal, modern society. It was ethnocultural and ethnoreligious impulses that dominated the voter’s political perspective during this preindustrial period.

Ethnocultural theories have been criticized for creating a false dichotomy between economic and cultural politics (Wilentz 1984). This critique seeks to understand the agrarian movement and labor radicalism of the late nineteenth century not only as cultural or social movements but also as programmatic third-party politics (Ritter 1997). In this view, the antimonopolist parties provided policy alternatives to farmers and laborers who were frustrated with both major parties’ handling of the nation’s finance. Ethnocultural theorists have held the view that economic programs of this era were related to voters more with the image and character of a party than with substantive policy proposals (McCormick 1974: 356–57). And Burnham (1965: 226–27) has emphasized the “profound antagonism in cultural and political style” in explaining the urban realignment in 1896, when the northern workmen produced a massive swing to the Republican Party (Sundquist 1983:

162–65).⁴ Still, it is quite possible that some form of economic voting might have occurred, at least within certain strata of the society.

Even so, the behavior would have been different from what we understand about economic voting today. The southern and western farmers, suffering from the enduring low price level in a time of constant economic boom-busts, would have voted not only against deflation but also for slower growth. Although in 1896 the agrarian insurgents largely remained loyal to the Democratic Party because of William Jennings Bryan's candidacy, in 1892 they supported James B. Weaver of the People's Party, who won 8.5% of the total vote. The overlap of populism with the Democratic Party allowed for the identification of the then incumbent Republican Party with corporate liberalism and industrial prosperity (Ritter 1997: 57).

The election of 1896 signified the start of the shift from "community" to "society" in terms of Hays's conceptual framework. It commenced a new political universe that Burnham finds fundamentally different from the preindustrial political system. The change was, of course, not abrupt but gradual. In 1928, Al Smith forged a Democratic majority in large cities on ethnocultural issues. But, ironically, as workingmen increasingly voted Democratic, socioeconomic voting was gaining momentum. As Hays concludes (1975: 160), the Great Depression and the New Deal crystallized the newly emerged "impulses arising from the technological organization of modern industrial society."

Today, although some political scientists are still pondering whether a realignment occurred during the 1960s (Aldrich and Niemi 1990), and some are even debating whether the notion of "realignment" still has much explanatory power (Shafer 1991), many would readily agree that the political universe of the New Deal party system has again been changed. Burnham (1991) uses the term *postindustrial/postparty politics* to distinguish the new era starting in the 1960s from the industrial order initiated in 1932. This coincides with what others call the "postmaterialist" era. Recent works by Ronald Inglehart (1990) and Ben Wattenberg (1995), for example, register a move from the secular, materialist perspectives of the cosmos toward values that are less mechanical and instrumental and that emphasize understanding the meaning and purpose of human life. Inglehart sees as likely a renewed concern for spiritual values—not a return to the agrarian society or a reprise

of traditional religion—which increases the importance of cultural factors as determinants of the shape of society and politics. The American voter, who had been alienated from the ethnocultural community during the process of industrialization, has gone a long way beyond socioeconomic concerns.

Interestingly, Inglehart (1990: 16) also criticizes what he calls “rational choice” models of economic voting for ignoring cultural variables. He claims that “with the emergence of advanced industrial society, the impact of economic factors reaches a point of diminishing returns” (*ibid.*: 285). As social and cultural issues such as abortion, homosexuality, and family values increasingly affect election outcomes, it is plausible that economic voting has reached a turning point in the postmaterialist political universe.

Ahistoricism in Time-Series Analysis of Voting

Aggregate-level analyses of economic voting are not totally ignorant of the changing historical context, but they tend to focus on the political circumstances rather than the social settings from which political behavior springs. Early works such as Kramer 1971, Stigler 1973, and Fair 1978 differ on how to deal with wartime elections and elections with a large third-party vote. George Stigler (1973) in particular points out the instability of Gerald Kramer’s results. He argues that 1900–1932 and 1934–70 should be examined separately because the Republicans dominated the first period and the Democrats the second, and his results show no statistical significance in the effect of economic conditions in congressional elections. Ray Fair’s (1978) model includes a trend term to account for some of the “unaccounted-for factors” that may have trends over time. These considerations notwithstanding, all three studies postulate a universal voter who looks at the national economy in making self-interested voting decisions. Their main theoretical difference is on the amount of information required of the voter (see Fair 1978: 160). Both Kramer’s and Stigler’s analyses of congressional elections start from 1896; Fair begins with 1916 in his analysis of presidential elections, saying that his equations do not fit data well prior to that year. While their results vary, it is apparent that they share the assumption that voters’ perspectives and behaviors are homogeneous throughout a long and fixed time frame.⁵

The ignorance of critical contingencies of social change in time-series

analyses of historical processes is criticized as “ahistoricism” by Isaac and Griffin (1989). Isaac and Griffin examine time-series studies in sociology, economics, and political science in a variety of substantive areas. They find (1989: 873) that almost every one of these studies ignores “the sudden or gradual temporal conditioning of historical-structural relationship.” They summarize the ahistorical practices they observe into nine “historically limiting premises,” including

the assumption that time-series coefficients are constant for the entire (and deliberately long) time-period over which the equation is estimated, thus a priori precluding the search for temporal conditioning of parameter estimates; (5) the practice, typically determined by data availability, of “slicing into” history at virtually any convenient point and ending at another convenient point even if, historically, the temporality of the process under examination differs from the happenstance of convenience. (1989: 878)

Griffin and Isaac (1992) further elaborate on the implications of the homogeneity assumption. Time-series analyses of historical events and processes premised on this assumption, Griffin and Isaac (1992: 167) conclude, “obscure temporally heterogeneous causal processes, conceal historical turning points in parameter structures, and thus miss or otherwise underestimate how historical contexts condition causality and how contingent actions establish possibilities for future action.” They advocate the use of recursive regression, a procedure adapted from the better-known “moving regression” technique (Brown et al. 1975), to assess the empirical validity of the homogeneity assumption.

In a study of the effect of political corruption on presidential elections, Tim Fackler and I (1995) make a conscious effort to pursue Burnham’s idea of comparative sociology and avoid Isaac and Griffin’s critique of ahistoricism. Citing the socioeconomic changes that led to the shift in Hays’s “community-society” continuum, Fackler and I argue that the conditions of voters’ processing of economic and political information, particularly information about corruption, was substantively changed by the New Deal. To substantiate our argument, we employ a measure of information about corruption and calculate its correlation with the presidential vote for the incumbent party over a sequence of short and moving time frames from 1890 to

1992. The analysis clearly indicates the time-varying nature of the effect of information about corruption in correspondence with the evolving historical context that conditions voting behavior. Although this study includes economic variables, its theoretical focus is corruption, and the historical effect of economic voting remains to be addressed.

Our theoretical and substantive understanding points to critical turns in the dynamic relationship between voting behavior and economic events. To avoid the critique of ahistoricism, the conventional homogeneity assumption concerning the effect of economic variables on election outcomes must be empirically tested. In other words, the stability of the relevant coefficients in a time-series model of economic voting must be examined. If evidence shows that these coefficients are unstable, then an appropriate model of economic voting must allow for the time-varying effect of economic events in changing historical contexts. In time-series analysis, techniques of time-varying parameter regression are readily available (Beck 1983). This article uses such techniques to examine the historical significance of economic voting.

In the rest of this article, I first introduce a model of economic voting primarily based on Fair 1988; Fair's periodic updates of his 1978 article constitute the main literature on economic conditions and the presidential vote at the aggregate level. I then conduct tests for the stability of the coefficients of this basic model. After presenting evidence of instability, I propose a time-varying parameter regression model for economic voting in historical contexts. A substantive discussion follows these empirical analyses.

The Basic Model and Evidence of Parameter Instability

The economic voting model examined is

$$V_t = \beta_1 + \beta_2 W_t + \beta_3 R_t + \beta_4 p_t I_t + \beta_5 g_t I_t + u_t \quad (1)$$

where $t = 1872, 1876, \dots, 1996$; V_t = the Democratic share of the two-party vote, in percent; $W_t = 1$ for $t = 1976$ and 0 otherwise; $R_t = 1$ if there is a Democratic incumbent and he is running for reelection, -1 if there is a Republican incumbent and he is running for reelection, and 0 otherwise; $I_t = 1$ if there is a Democratic incumbent, -1 if there is a Republican incumbent; p_t = the average annual inflation/deflation rate during a four-year presiden-

tial term; and g_t = annual growth rate of real per capital GNP in the election year. This model is similar to the equation estimated in Fair 1988, with important modifications, some, but not all, of which are consistent with Fair's 1992 update (1996a, 1996b).

First, W_t is included as an independent variable to capture the extraordinary effect of Watergate. This is necessary in the light of Fackler and Lin 1995, which demonstrates that information about corruption had a significant effect on presidential elections from 1932 to 1992 and whose measure for information about corruption shows the dominant role of Watergate. Fair does not explicitly include Watergate in his models, but his treatment of Ford as a nonincumbent (i.e., setting $R_t = 0$ for 1976) has the same effect. Such a practice is undesirable, however, because it raises the question of how to treat other running presidents who succeeded to office. With the inclusion of W_t , Ford is here considered as an incumbent.

Second, I_t is not included as a stand-alone independent variable, as it is in all versions of the Fair model. This is not only because it is never statistically significant but also because there is no particular reason to expect it to have a "main effect." I_t is multiplied by g_t and p_t to adjust the signs of their coefficients so that the effect of economic events is on the *incumbent* party's share of vote. Since both g_t and p_t are ratio scales, it is not necessary to include main effects unless there is a substantive or theoretical reason to do so (Cohen 1978).

Third, in the operationalization of p_t and g_t , I have followed Fair (1996a, 1996b) in extending the voter's retrospective horizon to a full presidential term for inflation/deflation and one year for growth rate. But note that annual, rather than quarterly, economic data are used for this study. Thus, the economic events in the fourth quarter of an election year are included in the calculation of p_t and g_t . Excluding fourth-quarter data leads to the exclusion of economic events in the critical month prior to election day, while including fourth-quarter data leads to the inclusion of postelection events. Unless monthly data are used, there is no good solution to this dilemma. Unfortunately, monthly data are not available for most historical periods.

Fourth, Fair's 1992 update drops an atheoretical (and statistically insignificant) linear trend term and adds in its place two new variables: the number of quarters of economic "good news" and a formula-based dummy variable representing the duration the incumbent party has been in power.

In Equation 1, the trend term is excluded, but the new variables are not included. Fair's goal is to increase forecastability in the wake of the model's disappointing performance in the 1992 election.⁶ Although including the new measures increases forecastability, the implementation is data-driven and entails additional parameters.⁷ Since the objective of this article is to test the stability of the effects of economic events rather than to improve forecastability, I opt not to include Fair's new variables.

Finally, Fair (1996a, 1996b) sets the effect of inflation to zero for 1920, 1944, and 1948 to account for the effect of wartime price control. In the following analysis, I do not explicitly constrain the coefficients for the war years. The time-varying parameter regression procedures used in estimating Equation 1 will reveal the effect of historical contingencies on the coefficients for both inflation/deflation and growth rate.

Historical data for all the variables in Equation 1 are available.⁸ A detailed description of the data and their sources is given in Appendix 1.

Figure 1a shows the plot of the growth rate of real per capita GNP from 1872 to 1996. Figure 1b shows the plot of the incumbent party share of the two-party presidential vote for the same period. Put together, the two time series seem to be following quite different paths: while growth fluctuated in a sequence of booms and busts from the beginning of the period to the time of the New Deal, electoral competition was kept close until after the turn of the century, when large partisan swings began to occur.⁹ The only historical period in which the two trajectories vary in an apparent synchronic fashion seems to be the three decades centered around the New Deal, suggesting that this could have been a period of significant economic voting.

In testing the conjecture that the significance of economic voting could differ in distinct historical periods, I first use a "moving Chow test" on Equation 1. The Chow test is often used to examine whether a regression model has different intercept and slope coefficients across two subperiods (or subsamples), while within each subperiod the parameters are assumed to be constant (Chow 1960). The test statistic is

$$F = \frac{[S_3 - (S_1 + S_2)]/k}{(S_1 + S_2)/(N_1 + N_2 - 2k)}$$

where k is the number of parameters in the regression model; N_1 and N_2 are the number of observations for the two subperiods; S_1 and S_2 are the sum

Figure 1 Time paths of economic growth and election outcomes

Table 1 F statistics for the moving chow tests

Periods compared	F statistics
1872–1908/1912–48	3.26 **
1876–1912/1916–52	12.00 ***
1880–1916/1920–56	10.80 ***
1884–1920/1924–60	1.93
1888–1924/1928–64	1.85
1892–1928/1932–68	1.36
1896–1932/1936–72	0.87
1900–1936/1940–76	0.56
1904–40/1944–80	0.55
1908–44/1948–84	1.99
1912–48/1952–88	1.24
1916–52/1956–92	1.43
1920–56/1960–96	1.93

** $p \leq .05$, *** $p \leq .001$; $df = 4, 20$.

of squared residuals for the two separate regressions estimated within each subperiod; and S_3 is the sum of squared residuals for the pooled regression estimated for the entire period (Gujarati 1988). The statistic follows the F distribution with degrees of freedom $df = k, N_1 + N_2 - 2k$. In implementing the test, a subperiod is taken to consist of 10 consecutive elections. First, the period 1872–1908 is compared with the period 1912–48. The dividing point is then shifted to 1912/1916, and so on. This “moving” approach makes it possible to examine the pattern of change over a continuous historical period.

Table 1 shows the results of the moving Chow test for presidential elections from 1872 to 1996. Interestingly, only for dividing points in the earlier years of the century is the test statistically significant at the .05 level. This finding is consistent with Fair’s observation that his equation does not fit the data well prior to 1916. It also supports the conjecture that economic events may not have the effects prescribed by the theory of economic voting until well into the “system of 1896.” The low F statistics at the beginning of the New Deal indicate a high level of structural homogeneity between the subperiods before and after, suggesting a historical context that is entirely consistent with economic voting. As the dividing point moves forward, the F values increase again, although they do not achieve clear statistical significance. The trend, however, is consistent with the postmaterialist contention

that the impact of economic factors might be reaching a point of diminishing returns.

Although the results seem satisfactory, it should be noted that the Chow test examines only structural stability, not the nature of the structural relationship in question. To further investigate the stability of the effect of economic events on voting, I run a moving regression analysis (Brown et al. 1975) on Equation 1.

The method is easily implemented by estimating the equation for a fixed subperiod starting from the first observation, assuming constant parameters. Then the estimation subperiod is shifted one observation forward, and so on, until the last observation is reached. The series of estimates of each parameter can then be plotted against time, as represented by the midpoint of the shifting subperiod. The plot provides informative insights about the stability (or lack thereof) of the parameters over time.¹⁰

Figure 2a shows the moving estimates of β_4 , the coefficient associated with inflation/deflation, and the corresponding 95% confidence limits.¹¹ The length of the subperiods used in these moving regressions is fixed at 15 elections, starting from 1872 and ending at 1996.¹² As expected, the effects (on votes for the incumbent party) are negative, but no statistical significance is clearly achieved except for a few estimates. The variation of the effects over time appears to be rather stable, except for a notable decrease in the World War II year 1944 and its aftermath, 1948. There is, however, no clear sign of wartime effect in 1916 or 1920, the World War I years.

The story is conspicuously different with respect to growth, as is illustrated in Figure 2b. Starting from initial insignificance at the midpoint 1900, the estimates turn from negative to positive, achieve statistical significance in the years before the New Deal, and remain significant for the rest of the estimation period. The magnitude of the estimates increases almost linearly from the beginning until about World War II, but afterward it appears stable.

The moving regression results suggest that, as far as parameter stability is concerned, inflation/deflation and economic growth should be treated differently. Reaction against inflation and deflation is probably of a more universal nature for all historical contexts, but society's sentiment toward economic growth is apparently more volatile and context-dependent. Thus, the structural instability evidenced by the moving Chow tests is mostly due to the parameter instability associated with growth, not to inflation/deflation.

Figure 2 Effect of economic events on votes for president: Moving regression estimates

Tracking the Historical Significance of Economic Voting

The moving regression method, however suggestive, does not provide a rigorous statistical test for parameter stability. The estimated effect at each point is actually the average effect for a fairly wide range of neighboring points. The method is also limited because no estimates are provided for the beginning and ending subperiods. This limitation especially thwarts our interest in examining parameter instability for the early and late twentieth century. Considering the moving regression method a reasonable first step in detecting parameter instability, time-series analysts have designed various models to add more structure to the process of parameter change. These models commonly assume that a regression parameter suspected of instability is an explicit function of time. They vary in the functional form and hence its substantive interpretation and estimation procedure. One class of such models uses a stochastic Markov functional form. For example, one may assume that β_5 in Equation 1, the effect of economic growth, follows a “random walk” scheme:

$$\beta_5(t) = \beta_5(t-1) + \varepsilon(t)$$

This model postulates that in any presidential election, the effect of economic growth on the vote changes from its previous value by a random error. Such a specification allows a parameter to drift upward or downward for a period of time. Lynn Roy LaMotte and Archer McWhorter (1978) design a test for regressions with random walk parameters. Of course, other stochastic Markov functional forms can be specified for an unstable parameter. Nathaniel Beck (1983) suggests estimation by Kalman filtering for this class of models. Because of the inclusion of a stochastic component in the functional form, estimated parameters usually show frequent short-term fluctuations.

Another class of time-varying parameter regression models simply assumes that the change in a parameter is smooth enough to be approximated by a polynomial function of time. In our economic voting model, specifying

$$\beta_5(t) = \sum_{j=0}^n \alpha_{5,j} t^j = \alpha_{5,0} + \alpha_{5,1} t + \alpha_{5,2} t^2 + \dots + \alpha_{5,n} t^n \quad (2)$$

with constant $\alpha_{5,j}$'s postulates a polynomial of the n th order for the changing effect of economic growth. In general, the minimum order of the polynomial can to some extent be determined by theory. For example, if we expect the effect of economic growth to increase in the early twentieth century and decrease later, the order should be $n = 2$ or higher. If, however, the expectation is that the effect first bottoms out before increasing and then saturates, then the order should be at least 3. Mathematically, each turnabout would require at least one additional order above the first-order polynomial, which prescribes a linear trend. Beyond the minimum order required by theory, a higher-order polynomial can exhibit similar long-term dynamics but will pick up short-term fluctuations as well. When the order of the polynomial cannot be completely determined theoretically, it must be justified by statistical procedures.

In what follows, I use a variant of this polynomial approach to track the historical significance of economic voting. The approach has the virtue that it is straightforward, flexible, and intuitively understandable. More important, our theoretical discussion in the previous sections has focused on long-term rather than short-term changes. A third-order polynomial model may be sufficient to track the expected long-term trends, while an equally parsimonious stochastic Markov model may obscure the global dynamics by tracking local dynamics too closely.

The specific model I use is the Legendre polynomial model developed by Melvin Hinich and Richard Roll (1981). Hinich and Roll suggest using the mutually orthogonal Legendre polynomials $P_j(z)$ instead of t^j in modeling a time-varying parameter:

$$\beta_5(t) = \sum_{j=0}^n \beta_{5,j} P_j(z)$$

where z is rescaled time and $\beta_{5,j}$ ($j = 1, 2, \dots, n$) are constants to be estimated. Technical details and the estimation procedures for this model are provided in Appendix 2. It suffices to say here that this model is mathematically equivalent to the usual polynomial model (2) but lacks the undesirable collinearity among the powers of time: $t, t^2, \dots, t_n$.

In implementing this model to Equation 1, I assume that β_1, β_2 , and β_3 (i.e., parameters associated with, respectively, the intercept, Watergate

dummy, and incumbency) are fixed over time, as there is no theoretical reason to expect otherwise. Furthermore, the moving regression results suggest that β_4 (the effect of inflation/deflation) is likely to be stable, except perhaps for World War II years, and β_5 (the effect of economic growth) is likely to be unstable. Such information is useful in modeling β_4 and β_5 .

Table 2 presents a third-order polynomial solution for β_5 in conjunction with a zero-order solution for β_4 . Note that the estimate of $\beta_{5,3}$, like the estimate of $\beta_{4,0}$, is statistically significant. No other models of all possible combinations of polynomials for β_4 and β_5 up to the sixth order have a statistically significant order higher than zero for β_4 and three for β_5 . Indeed, no other models in all these 36 combinations have a smaller standard error of regression (3.23) than this model, which improves substantially from the fixed-parameter “naive” model (SER = 4.71), also shown in Table 2. Thus, before venturing into even higher orders, I settle for this model, which implies stability in the effect of inflation/deflation but instability of the third order in the effect of growth. Setting $\beta_{4,0}$, the coefficient for inflation/deflation, to zero for 1944 and 1948 does not significantly change the estimates. Figure 3a shows the time path and the 95% confidence limits of the effect of economic growth represented by the third-order solution.

The third-order model for the effect of growth on voting is strikingly consistent with our theoretical expectations: an upturn at the onset of industrialization and a downturn during the postmaterialist cultural shift. The lack of positive effect in the late nineteenth century apparently reflects the dominance of ethnocultural concerns. While ethnocultural theories do not exactly anticipate the weakly negative, yet statistically significant, effect centered around 1892, such negativity is arguably related to the agrarian revolt and the rise of populism that began in the 1870s.

From a western farmer’s perspective, growth benefited only the eastern monopolies and speculators, while slicing prices for their products (Sundquist 1983: 108–9). As the machine-dominated Democratic and Republican Parties rejected reforms, the antimonopolists had no choice but to turn to third-party movements (Ritter 1997: chap. 2). The antisystem nature of these movements imputed certain disadvantages on the incumbent parties throughout most of this period (see Figure 1b). It is not coincidental that in 1892 General James B. Weaver and his People’s Party won more than 8% of the total vote, including a plurality in five states.

Table 2 Hinich-Roll time-varying parameter regression results

Dependent variable: % Democratic share of the two-party vote				
Variable parameter	1872-1996 OLS estimates ^a	1872-1996 GLS estimates ^b	1872-1988 GLS estimates ^b	1872-1996 GLS Estimates ^b
Constant	β_1 49.171 (57.37)***	47.794 (64.55)***	47.173 (76.77)***	46.921 (53.56)***
Watergate	β_2 3.539 (0.71)	8.969 (2.49)*	10.043 (3.42)**	10.514 (2.01)+
Incumbency	β_3 5.359 (3.65)**	4.115 (3.92)***	4.178 (4.05)***	4.333 (3.52)***
Inflation/deflation	$\beta_{4,0}$ -.665 (-2.81)**	-.687 (-3.73)**	-.615 (-4.09)***	-.454 (-2.14)*
Economic growth	$\beta_{5,0}$.268 (1.91)+	.550 (4.09)***	.528 (5.24)***	78.900 (2.08)+
	$\beta_{5,1}$	1.116 (4.12)***	1.138 (5.48)***	.765 (2.36)*
	$\beta_{5,2}$	-.381 (-1.13)	-.478 (-1.64)	155.745 (2.06)+
	$\beta_{5,3}$	-1.052 (-3.47)**	-1.109 (-4.25)***	-1.567 (-2.98)**
	$\beta_{5,4}$			-.742 (-1.36)
	$\beta_{5,5}$			-.217 (-.34)
	$\beta_{5,6}$			.003 (.00)
	$\beta_{5,7}$			1.097 (1.93)+
	$\beta_{5,8}$			.574 (.87)
	$\beta_{5,9}$			.492 (.83)
	$\beta_{5,10}$			1.347 (2.02)+
	$\beta_{5,11}$			1.043 (1.36)
	$\beta_{5,12}$			1.489 (2.07)+
N	32	32	30	32
R ²	.44	.74	.83	.78
SER	4.71	3.23	2.60	2.97

+ $p \leq .10$ * $p \leq .05$ ** $p \leq .01$ *** $p \leq .001$, two-tailed test; t statistics in parentheses.

^aNo evidence for AR(1) autocorrelation ($d = 1.88$).

^bThe Durbin-Watson test for AR(1) autocorrelation is inconclusive, but GLS was used because of the low degrees of freedom. See Johnston 1984: 316.

Figure 3 Effect of economic growth on votes for president: Hinich-Roll time-varying parameter regression estimates

The negative effect of growth on incumbent votes is, in a sense, a kind of “economic voting” that is alien to the modern notion of the phrase. It does not negate ethnoculturalists’ theory of voting, but it does require that what Gretchen Ritter (1997) calls “the politics of finance” of this period be taken more seriously. In any case, it reinforces the contention that voting behavior must be examined within its historical contexts.

After turning positive, the effect quickly achieves statistical significance before the New Deal and keeps rising through the war years until it saturates in the social upheavals of the 1960s. As racial and gender politics and the Vietnam War play their way into electoral politics, the effect starts declining. By 1988, what remains of the positive effect has become statistically insignificant, as the confidence band increases. It seems convenient, then, to say that the cultural shift anticipated by postmaterialist theorists has manifested itself in the decline of economic voting, a trend that continues to the present. After all, abortion, the gender gap, homosexuality, and family values were the dominant issues in recent elections. And the failure of the Fair model in forecasting the 1992 election seems only to confirm the dwindling significance of economic voting.

A puzzling hitch in such reasoning is this: if the effect of growth on voting has indeed declined in recent elections, then the fit of the third-order solution (for the effect of economic growth) should be as good as the same model estimated a few elections ago, if not improving. This is not the case, however. As shown in Table 2, the same model estimated for the 1872–1988 period sports a much smaller SER (2.60). But if the fit of the model has become worse, does this imply that the effect of growth has not been following the prescribed course?

To answer this question, I increment the order of Legendre polynomials for β_5 while leaving all other parameters fixed. Following the procedure explained in Appendix 2, I arrive at a 12th-order solution, where the last parameter is weakly significant.¹³ Estimated for the 1872–1996 period, this model has an improved SER (2.97; see Table 2), so it’s arguably a better model. The estimated time-varying effect of growth and its confidence limits are shown in Figure 3b.¹⁴

The plot clearly demonstrates that the 1992 election is a singular case. The trajectory of the effect otherwise generally parallels that of the third-order solution, and more detailed local texture provides additional informa-

tion. The 1876 “midsequence crisis” (Burnham 1991: 118–19) now exhibits an amplified negative effect, although still insignificant. The effects at 1892 and 1896 remain negatively significant. And then an extended period of almost linear increase follows until the World War II years clearly switch the voters’ focus. After the war years the trend resumes, rising and peaking in the 1960s just as the third-order model shows. The ensuing postmaterialist decline is also unambiguous. But the dramatic turn at 1992 into negative, albeit barely significant, is totally beyond expectation.¹⁵

Since it is hardly likely that a 1.5% growth in per capita GNP could turn into a loss of more than five points for George Bush in 1992, the oddity seems to be related to Ross Perot’s candidacy. As documented in Appendix 1, the dependent variable in this study excludes third-party votes except for 1912 and 1924. Although Perot won up to 18.9% of the total votes, exit polls showed that his support came about equally from voters who would have voted for Bush or Bill Clinton without Perot’s candidacy.¹⁶ Thus, the two-party vote should have resembled the outcome of a hypothetical two-candidate race. Still, Fair reports that when, on the suggestion of Everett Ladd (1993), he pools together Bush and Perot votes, his equation no longer shows a large prediction error for 1992 (Fair 1996b: 123). In this study, such a regrouping would produce a third-order model for 1872–1996 with an SER of 2.57, slightly better than the same model estimated for the 1872–1988 period! Third-party movements usually attract voters who have deep distrust in the major-party system. As evidenced by the antimonopolists in the late nineteenth century, these voters can usher in a totally different pattern of economic voting. It is thus not unreasonable to suggest that Perot’s candidacy caused the singularity of 1992 in economic voting. Because of 1992, aggregate-level analysis of economic voting has lost some of its credibility (Greene 1993). It is remarkable that the time-varying parameter regression approach is able to identify the singularity of the 1992 election.

Conclusion

More than three decades have passed since the publication of Hays’s “The Social Analysis of American Political History, 1880–1920” (1965), and Burnham’s “The Changing Shape of the American Political Universe” (1965). As Burnham anticipated, it has become quite common, even imperative,

for researchers to relate aggregate data to findings of survey research. Because of the advantage of survey research in discovering the psychological and socioeconomic forces driving individual behavior, great scientific advances have been made through scholarly efforts to link the “macro” to the “micro.” Burnham’s warning remains valid, however: findings of survey research about the contemporary American electorate are not necessarily relevant to earlier periods of our political history. For historical periods, as Hays suggested, a social analysis of the circumstances in which people then lived and the ways in which they expressed their goals and values remains essential to aggregate analyses.

As aggregate data are accumulated for an ever growing and ever changing time period, it has become even more necessary for analysts to attend to the issues that Hays and Burnham raised. This article has demonstrated that the effect of economic voting changes with the changing historical context. By ignoring historical contingencies, the conventional aggregate analyses of economic voting open themselves up to the criticism that they are ahistorical.

Appendix 1: Data Sources

Sources

1. Balke and Gordon: Nathan S. Balke and Robert J. Gordon (1986) “Appendix B: Historical data,” in Robert J. Gordon (ed.) *The American Business Cycle: Continuity and Change*. Chicago: University of Chicago Press.
2. NIPA-86: *The National Income and Product Accounts of the United States, 1929–1982* (1986) Washington: United States Department of Commerce.
3. NIPA-97: United States Department of Commerce (1997) “Summary National Income and Product Series, 1929–1996.” *Survey of Current Business* (May).
4. HSUS: *Historical Statistics of the United States, Colonial Times to 1970* (1976) Washington: United States Bureau of the Census.
5. CQ Guide: *Congressional Quarterly’s Guide to U.S. Elections* (1985), 2d ed. Washington: Congressional Quarterly.
6. CQ Report: *Congressional Quarterly Weekly Report*, various issues.
7. EROP: *Economic Report of the President* (1997) Washington: United States Government Printing Office.

Data

t : 1872, 1876, . . . , 1996.

V_t : Democratic share (%) of the two-party vote.

1872–1984: CQ Guide.

1988–96: CQ Report.

Except for 1912 and 1924, only the official candidates of the two major parties are included in the calculation. For 1912, William Howard Taft and Theodore Roosevelt are pooled to calculate the Republican vote. For 1924, John W. Davis and Robert M. La Follette are pooled to calculate the Democratic vote.

GNP_{1_t}/GNP_{2_t}: GNP (in billions of current/constant dollars).

1869–1928: Balke and Gordon.

1929–96: NIPA-97.

POP_t: U.S. population (millions).

1869–1928: HSUS.

1929–58: NIPA-86.

1959–96: EROP.

g_t : Annual growth rate of real per capita GNP.

$$g_t = \frac{(\text{GNP}2_t / \text{POP}_t) - (\text{GNP}2_{t-1} / \text{POP}_{t-1})}{(\text{GNP}2_{t-1} / \text{POP}_{t-1})} \times 100\%$$

p'_t : Annual growth rate of the GNP implicit price deflator.

$$p'_t = \frac{(\text{GNP}1_t / \text{GNP}2_t) - (\text{GNP}1_{t-1} / \text{GNP}2_{t-1})}{(\text{GNP}1_{t-1} / \text{GNP}2_{t-1})} \times 100\%$$

p_t : Average annual inflation/deflation rate during a presidential term.

$$p_t = |(p'_t + p'_{t-1} + p'_{t-2} + p'_{t-3})| / 4.$$

Appendix 2: Hinich-Roll Time-Varying Parameter Regression

For time-series regression

$$Y_t = \beta_1 + \beta_2 X_{2t} + \dots + \beta_k X_{kt} + u_t \quad t = 1, 2, \dots, T$$

Hinich and Roll (1981) suggest that a time-varying parameter, β_i , can be specified as

$$\beta_i(t) = \sum_{j=0}^n \beta_{i,j} P_j(z(t)) \quad (3)$$

where $P_j(z(t))$ are the Legendre polynomials (see Table 3); n is the highest order of the polynomials; and $\beta_{i,j}$ ($j = 1, 2, \dots, n$) are fixed parameters. The compelling reason for using the Legendre polynomials rather than the usual powers of time in this specification is that the former are mutually orthogonal in the interval $[-1, 1]$ and hence without the high collinearity among the latter. To ensure orthogonality, however, t ($t = 1, 2, \dots, T$) has to be rescaled into $z(t) = [2(t - 1)/(T - 1)] - 1 \in [-1, 1]$. For example, for the historical period (1872–1996 with $T = 32$) used to estimate Equation 1, $z(t) = (2/31)(t - 1) - 1$.

Table 3 Legendre Orthogonal Polynomials

$$P_0(z) = 1$$

$$P_1(z) = z$$

$$P_2(z) = (3z^2 - 1)/2$$

$$P_3(z) = (5z^3 - 3z)/2$$

$$P_4(z) = (35z^4 - 30z^2 + 3)/8$$

$$P_5(z) = (63z^5 - 70z^3 + 15z)/8$$

$$P_6(z) = (231z^6 - 315z^4 + 105z^2 - 5)/16$$

$$P_7(z) = (429z^7 - 693z^5 + 315z^3 - 35z)/16$$

$$P_8(z) = (6,435z^8 - 12,012z^6 + 6,930z^4 - 1,230z^2 + 35)/128$$

Source: Calculated from Rodrigues's Formula: $P_n(z) = (-1)^n / (2^n n!) (d^n / dz^n) \{(1 - z^2)^n\}$. See Beyer 1984: 372.

With β_i modeled as such, the term $\beta_i X_{it}$ in the regression becomes

$$\beta_i X_{it} = \left[\sum_{j=0}^n \beta_{i,j} P_j(z) \right] X_{it} = \sum_{j=0}^n \beta_{i,j} [P_j(z) X_{it}]$$

Since $P_j(z)X_{it}$ can be evaluated with empirical data, the regression is still linear in the parameters ($\beta_{i,j}$) and hence can be estimated by conventional methods.

To illustrate, suppose β_i is a third-order polynomial of time. Instead of using the conventional expression

$$\beta_i(t) = \alpha_{i,0} + \alpha_{i,1}t + \alpha_{i,2}t^2 + \alpha_{i,3}t^3$$

and estimating the $\alpha_{i,j}$'s as the coefficients of $t^j X_{it}$ ($j = 0, 1, 2, 3$), the Hinich-Roll method uses

$$\beta_i(t) = \beta_{i,0}P_0(z) + \beta_{i,1}P_1(z) + \beta_{i,2}P_2(z) + \beta_{i,3}P_3(z)$$

and estimates the $\beta_{i,j}$'s as the coefficients of $P_j(z)X_{it}$. Substituting $P_j(z)$ with the appropriate polynomials from Table 3 and z with $(2/31)(t - 1) - 1$ gets

$$\begin{aligned} \beta_i(t) = & (\beta_{i,0} - 1.0645\beta_{i,1} + 1.1998\beta_{i,2} - 1.4190\beta_{i,3}) \\ & + (.0645\beta_{i,1} - .2060\beta_{i,2} + .4515\beta_{i,3})t + (.0062\beta_{i,2} - .0332\beta_{i,3})t^2 + (.0007\beta_{i,3})t^3 \end{aligned}$$

which is still a third-order polynomial of t . The approach is thus no more than a reparameterization of the $\alpha_{i,j}$'s into the $\beta_{i,j}$'s. Like the $\alpha_{i,j}$'s, the $\beta_{i,j}$'s define the shape of the polynomial and are interpretable to the extent that the $\alpha_{i,j}$'s are.

Although it is preferable to justify the choice of n on theoretical grounds, in practice it is difficult to do so because theories may not be specific enough and different polynomials may be equally justified. To further complicate the matter, although a lower-order polynomial has the merit of parsimony, a higher-order polynomial may fit the data better. Hinich and Roll, admitting that the optimal number of higher-order polynomials cannot be determined objectively, suggest the composite hypothesis test for alternative n :

$$H_0: \beta_{i,1} = \beta_{i,2} = \dots = \beta_{i,n} = 0 \quad (4)$$

i.e., using an F test to determine whether all except the zero-order coefficient are not significantly different from zero. If H_0 can be rejected at a given significance level, one can conclude that β_i “has actually varied in some way,” and hence instability can be inferred (Hinich and Roll 1981).

This approach is questionable because while it does test stability, it does not justify the superiority of n over $n - 1$ except in the special case when $n = 1$. For example, a 12th-order polynomial may satisfy the test in 4, but it is possible that the estimate of $\beta_{i,12}$ may not be significantly different from zero and that a polynomial of the 11th order or lower may produce a better fit. A more sensible approach is to test n against $n - 1$ —that is, to require the last coefficient to be statistically significant by the conventional t test.¹⁷ Thus, the order of instability in a parameter can be determined by incrementing the order of Legendre polynomials.¹⁸ This is the approach adopted in this article.

When the effects of more than one independent variables in a regression are suspected of instability, fitting polynomial models to all the relevant β_i ’s could be difficult. Although the Legendre polynomials are orthogonal when time is rescaled to $[-1, 1]$, multicollinearity can still cause problems when the same polynomial $P_j(z)$ is multiplied by two or more variables. Very often incrementing n for one variable changes the pattern of statistical significance in another. When this happens, a blanket search of all possible combinations may be useful. Still, it is preferable to have some prior expectations about the order of instability in each variable.

Because of parameter instability, the statistical significance of the effect of growth on votes for president has to be examined throughout the historical period of interest. For an n th-order Legendre polynomial solution, the estimated effect

$$\hat{\beta}_i(t) = \sum_{j=0}^n \hat{\beta}_{ij} P_j(z(t))$$

is a function of time and can be plotted to show the trajectory of the estimated effect over the course of history. The variance of the effect at any point of time is thus

$$\text{Var}(\hat{\beta}_i(t)) = \sum_{j=0}^n P_j(z(t))^2 \text{Var}(\hat{\beta}_{ij}) + 2 \sum_{0 \leq h < k \leq n} P_h(z(t)) P_k(z(t)) \text{Cov}(\hat{\beta}_{ih}, \hat{\beta}_{ik})$$

where $\text{Var}(\hat{\beta}_{ij})$ and $\text{Cov}(\hat{\beta}_{ih}, \hat{\beta}_{ik})$ can be estimated by the sample variances and covariances of the $\hat{\beta}_{ij}$ ’s.

Notes

Tse-min Lin is an associate professor of government at the University of Texas at Austin. His research has focused on electoral studies in general and the dynamics of American elections in particular. He wishes to thank John Freeman, Jim Granato, David Prindle, Gretchen Ritter, Daron Shaw, and the editor and anonymous referees of *Social Science History* for their helpful comments and suggestions on early drafts of this article.

- 1 Availability refers to whether an attitude is stored in memory, while accessibility refers to the readiness with which a stored attitude is retrieved from memory (Al-drich et al. 1989: 125).
- 2 For positive findings on economic voting in either presidential or congressional elections based on aggregate data, see Kramer 1971; Tufte 1975, 1978; Bloom and Price 1975; Goodman and Kramer 1975; Fair 1978, 1982, 1988, 1996a, 1996b; Hibbs 1987; Erikson 1989; Alesina et al. 1993. For negative findings on congressional elections, see Stigler 1973; Arcelus and Meltzer 1975a, 1975b; Erikson 1990.
- 3 Individual-level evidence of economic voting is stronger for presidential elections than for congressional elections. In general, scholars have found that both personal and national economic conditions influence the presidential vote, but there is scant evidence for the effect of personal economic conditions on the congressional vote (Fiorina 1978; Kinder and Kiewiet 1979, 1981; Markus 1982).
- 4 According to Burnham (1965: 26), William Jennings Bryan, the archetype of "the parochial, dry, Anglo-Saxon Protestant inhabitants of rural areas," could have been nothing but the target of ethnocultural hostility by those belonging to the "cosmopolitan, immigrant, wet, largely non-Protestant components of American urban population."
- 5 Some aggregate studies of economic voting focus only on the postwar period. Edward Tufte (1975, 1978) and Robert Erikson (1989) confine their analyses to 1948 onward primarily because they use variables from survey data that were not available for earlier elections. Douglas Hibbs (1987), focusing on 1952–84, provides an exceptional discussion on how the structure of the American political economy changed profoundly from the prewar to the postwar period.
- 6 In an interview with Elizabeth Kolbert of the *New York Times* after the 1992 election, Fair said, "Perhaps the model could be adjusted in a way that would preserve its utility." Or perhaps, he conceded, "you just have to dump it." See Kolbert 1992.
- 7 "Good news" is defined as the number of quarters of the first 15 quarters of each presidential term in which the growth rate is greater than r percent. The coefficient of "good news" is set to zero for 1920, 1944, and 1948. "Duration" is zero if the incumbent party has been in power for one or two consecutive terms; it is $1 + (d - 3)k[-(1 + (d - 3)k)]$ if the incumbent party has been in power for $d \geq 3$ terms. Thus, in addition to the new coefficients associated with these terms, r and k also have to be estimated. Fair uses a grid search to find that $r = 2.9$ and $k = .25$ produce the best fit, though he does not provide the standard errors of these estimates (Fair 1996a: 125, 127). Not counting r and k , Fair's 1996 model involves eight parameters, to be estimated with a total of 20 observations. The moving regression procedure used later in this section estimates an equation for only 15 consecutive elections; hence, it is desirable to avoid overparameterization.
- 8 For historical economic data, I use the annual data compiled by Nathan Balke and Robert Gordon (1986), which go back as far as 1869. Fair (1996a, 1996b) uses quarterly data from the same source, which go back only to 1875. Christina Romer (1989)

provides an alternative set of estimates. Using Romer's data does not change the substantive conclusion reached in this article.

- 9 The Dicky-Fuller unit root test shows that both time series are stationary.
- 10 The moving regression method used here was designed by R. L. Brown, J. Durbin, and J. M. Evans (1975) to detect continuous parameter drift in time-series regressions. Based on a fixed number m of continuous time periods, the method estimates moving regressions on subsets of data 1 through m , 2 through $m + 1$, . . . , and $T - m + 1$ through T , where T is the total number of time periods in the sample. Griffin and Issac (1992) advocate a variant of this moving regression method. Their variant, called recursive regression, employs three separate approaches: forward, backward, and diagonal. The diagonal approach estimates the Brown-Durbin-Evans moving regressions. It is suggested, however, that these results can be buttressed by forward recursive regressions, which are estimated on subperiods 1 through m , 1 through $m + 1$, . . . , and 1 through T , and backward recursive regressions, which are estimated on subperiods $T - m + 1$ through T , $T - m$ through T , . . . , and 1 through T . The diagonal approach alone is limited, Griffin and Issac (1992: 170) contend, because "the extent of the estimated change or consistency in the parameter structures is likely a function of the historical events and actions 'captured' by the fixed number of time points defining the time frame."

While it is true that moving regression results can sometimes be affected by the choice of the time frame (i.e., m), the advantage of having a fixed time frame is not only methodological (Griffin and Issac 1992: 170) but also substantive. A fixed time frame defines a given historical context, i.e., the time periods that surround a historical moment and can throw light on our understanding of it. By using a constant time frame, moving regressions allow us to compare the contextual dynamics affecting different historical moments. In contrast, because they increment or decrement a time frame with one fixed end, forward and backward recursive regressions cannot claim to capture historical contextuality in a systematic way. The issue of sensitivity to the choice of time frame can be addressed to some extent by trying different values of m (see note 12). More important, the moving regression method, while quite informative, is but a first step in detecting parameter instability (Beck 1983). With a more structured approach, such as the Hinich-Roll polynomial model used in the next section, the issue of time frame becomes moot because an estimate for a time-varying parameter is provided for each and every historical moment. It is for these reasons that I choose to use moving (or diagonal) regressions but not forward and backward regressions in this article.

- 11 All the moving regressions are corrected for AR(1) autocorrelation, although the Durbin-Watson tests are mostly inconclusive due to the low degrees of freedom. Jack Johnston (1984: 316) justified this conservative practice by arguing that "the consequences of accepting H_0 when autocorrelation is present are almost certainly more serious than the consequences of incorrectly assuming it to be absent."

- 12 Experimenting with subperiods of smaller length produces essentially similar results.
- 13 Note that the 7th- and 10th-order coefficients become significant ($p \leq .10$) only after the 12th-order coefficient is added.
- 14 Once again, constraining $\beta_{4,0}$ to zero for 1944 and 1948 does not qualitatively change the results.
- 15 The negative effect at 1988 and the large positive effect at 1996, both of which remain statistically insignificant, are probably affected by the singularity at 1992.
- 16 Perot also pulled about equally from the major-party candidates in 1996, according to statewide tracking polls done for the Robert Dole–Jack Kemp campaign by Tony Fabrizio.
- 17 When no coefficient beyond a certain order m is significant, then the F test may be applied to test n against m when necessary. Such a test would determine whether the order of instability is greater than m . In this sense, the Hinich–Roll approach is just a special case of the approach recommended here.
- 18 Alternatively, it can be determined by *decrementing* from a sufficiently large order until the last coefficient becomes statistically significant.

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